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Money, Growth and Real Business Cycles: Some International Evidence

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Abstract

This paper addresses three questions commonly identified with the real business cycle approach: do nominal variables Granger-cause real outputs?; do permanent innovations account for a substantial portion of variability in output growth?; do the transitory disturbances reflect a small set of common aggregate shocks, or are there many diverse shocks? These questions are analyzed in the context of a statistical model whose state vector parallels that in the theoretical model of Long and Plosser (1982), and which explicitly considers the existence of both deterministic and common stochastic trends. Monthly data on sectoral industrial production from six countries suggest a strong causal role for money growth. However, such causality is not inconsistent with a direct test of the "reverse causation" hypothesis. Permanent innovations account for between 20 to 60 percent of the monthly variability in sectoral output growth, while the transitory innovations are only weakly contemporaneously correlated across sectors.

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1. Introduction

Real business cycle theory interprets cyclical economic fluctuations as the competitive, equilibrium outcome of agents' responses to changes in endowments, tastes and technology (Kydland and Prescott (1982), Long and Plosser (1983)). Broadly interpreted, this line of research has focused on two themes: one concerns the nature of impulse mechanisms and the other concerns the form of propagation mechanisms.

Since real business cycle (RBC) models typically abstract from informational rigidities or price inertia, monetary impulses are not singled out as the dominant source of economic fluctuations. Any observed co-movement between monetary variables and real quantities instead arises from reverse causation, or from the endogenous optimizing response of a banking sector providing transaction services that in turn serve as produced inputs to other sectors (King and Plosser (1984)). Rather, technological shocks or changes in government spending and tax policies are candidates for business cycle impulses.

Such impulses may reflect the influence of many uncorrelated, disaggregate factors (Long and Plosser (1987)). Agents' desires to smooth consumption, as well as the nature of produced inputs, however, serve to propagate the impulses throughout the system, leading to persistence and co-movement of endogenous variables. Further insight into the nature of these propagation mechanisms may be obtained by requiring that they be compatible with endogenous, sustained economic growth (King and Rebelo (1986)). Accordingly, recent research by King, Plosser and Rebelo (1987) and Prescott (1986) view the economy

as evolving along a non-deterministic equilibrium growth path, which is itself buffeted by disturbances. A key result that emerges is that it is not straightforward to isolate "short-run" transitory cyclical fluctuations around the growth path from "long-run" permanent variations in the growth trend itself (Nelson and Plosser (1982)).

Previous empirical tests of RBC models have examined the evidence on impulse and propagation mechanisms, more or less as separate issues. Tests of the importance of monetary disturbances have been provided by Bernanke (1986), Eichenbaum and Singleton (1986). Evidence on the importance, and nature, of propagation mechanisms appears in King and Plosser (1987), King, Plosser, Stock and Watson (1987). As King (1986) notes in a perspicuous discussion, however, the two themes cannot be treated in isolation for the purposes of testing RBC models. In particular, meaningful tests of the relative importance of monetary versus real impulses requires careful delineation of the pertinent state variables as well as the dynamic propagation mechanisms describing the evolution of these variables. For example, if output shocks are serially correlated, then standard methods of correcting for such autocorrelation would create the false impression that lagged money growth helps to predict future output. More generally, omitting relevant state variables would lead to erroneous rejections of noncausality. Sims (1980a,b) interprets the interest rate as a proxy for such state variables, and finds that the explanatory power of money growth is reduced when the interest rate is included in the system. Tests of money-income causality also appear to be sensitive to methods for treating

deterministic and stochastic trends (Eichenbaum and Singleton (1986), Christiano and Ljungqvist (1987), Stock and Watson (1987a)).

This paper attempts a unified empirical examination of impulse and propagation mechanisms in the context of a stylized RBC model along the lines of Long and Plosser (1983). As in that model, a vector of sectoral commodity outputs serves as the relevant state variable. The propagation mechanisms are isolated in the context of a log-linear model which allows for common stochastic trends arising from permanent shocks, and also for transitory shocks. Using recent advances in time series econometrics (Engle and Granger (1987), Sims, Stock and Watson (1986), Stock and Watson (1987b)), the estimated model can be used to address three questions central to RBC analysis: do nominal variables Granger-cause real outputs?; how much of the variability in output growth can be attributed to permanent innovations?; do the transitory disturbances in the model reflect a small set of common aggregate shocks, or are there many diverse shocks? Insofar as the specification of both the state vector and of the propagation mechanisms are compatible with stylized RBC models, sharper tests of the hypotheses underlying such models may be obtained.

The statistical model used to address these questions is formulated in the next section. Subsequent sections of the paper apply the model to monthly time series data from six industrialized nations. The experience of other countries, such as Japan and West Germany, to the extent that they display differences in the level and variability of secular growth rates, may serve to highlight the distinction between permanent and transitory innovations. A description of the data

set is provided in section 3, which also develops the specification of the model. Section 4 tests for money-income causality, and also evaluates an explanation for causality in terms of reverse causation. Evidence is presented in section 5 on the explanatory power of permanent innovations, and also on the commonality of transitory shocks (within a given country). A final section contains a summary and the conclusions.

2. Statistical Model

A simple statistical model that captures the key elements of RBC analysis is given by (King, Plosser, Stock and Watson (1987)):

$$x_t = \delta + H\tau_t + B(L)a_t$$

$$\tau_t = \mu + \tau_{t-1} + \eta_t. \qquad (1)$$

x_t is an n-dimensional vector of sectoral outputs. δ is a vector of constants. The behavior of x_t is influenced both by a k-element ($k \leq n$) nonstationary component, τ_t, through the constant matrix H, and by a n-element stationary component, a_t, through the matrix polynomial B(L) in the lag operator L: $B(L) = \sum_{j=0}^{\infty} B_j L^j$. It is assumed that the nxk matrix H has full column rank, and that the components of the sequence of matrices $\{B_j\}$ are absolutely summable. The nonstationary component evolves as a random walk with drift μ and innovations η_t where $E[\eta_t \eta_t'] = I_k$, the identity matrix of order k. The innovations η_t represent permanent shocks: a change in η_t has a once-and-for-all effect on the level of x_t which is not subsequently

reversed. On the other hand, a_t is a mean zero, serially uncorrelated disturbance vector with covariance matrix Σ, representing transitory, mean-reverting shocks.

System (1) serves as a simple dynamic reduced-form specification of the propagation mechanisms arising from an equilibrium growth model of the type analyzed by King and Rebelo (1986), King, Plosser and Rebelo (1987). The individual elements of the state vector move together because they share a smaller, common set of nonstationary stochastic trends. Innovations to these trends alter the steady-state growth characteristics of the model economy. In addition, transitory disturbances may also temporarily move the economy away from the trend.

Neither the common trends nor the transitory disturbances are observable. Accordingly, for purposes of estimation and testing, it is simpler to work with an alternative parameterization: first in terms of the fundamental moving average representation and, ultimately, in terms of the relationship between current and lagged values of the observed variables x.

Suppose that the fundamental Wold moving average representation of the first differences of x_t is given by:

$$(1-L)x_t = \gamma + C(L)\varepsilon_t. \qquad (2)$$

The model (1), together with the content of RBC analysis, generate the following restrictions on the matrix polynomial C(L). Multiply both sides of (1) by the matrix $(I_n - H(H'H)^{-1}H')$ to obtain

$$(I_n - H(H'H)^{-1}H')x_t = (I_n - H(H'H)^{-1}H')\delta + (I_n - H(H'H)^{-1}H')B(L)a_t, \qquad (3)$$

where I_n is the nxn identity matrix. The matrix $(I_n - H(H'H)^{-1}H')$, which has rank n-k, transforms x_t to a stationary process given by the right hand side of (3). In the terminology of Engle and Granger (1987), x_t is co-integrated with cointegrating rank n-k. Put another way, the presence of a set of k common stochastic trends implies that C(1) in (2) has rank k, so that conventional methods of fitting vector autoregressions are inappropriate. Further, suppose that the state vector x_t is augmented by a vector y_t of nominal variables (including, for example, the money supply). Assuming y_t has a Wold representation in first differences, the augmented system is[1]

$$\begin{pmatrix} (1-L)x_t \\ (1-L)y_t \end{pmatrix} = \begin{pmatrix} \gamma_x \\ \gamma_y \end{pmatrix} + \begin{pmatrix} C_{xx}(L) & C_{xy}(L) \\ C_{yx}(L) & C_{yy}(L) \end{pmatrix} \begin{pmatrix} \varepsilon_{xt} \\ \varepsilon_{yt} \end{pmatrix}. \qquad (4)$$

If, given the past history of x, the behavior of past nominal variables does not influence the future behavior of x, then $C_{xy}(L) = 0$.[2]

The Granger representation theorem (Engle and Granger, 1987) implies that the co-integrated system (2) may be re-written as the vector error-correction mechanism

$$(1-L)x_t = \phi + A(L)(1-L)x_{t-1} + d(\alpha' x_{t-1}) + \varepsilon_t. \qquad (5)$$

ϕ is a vector of constants.[3] A(L) is a matrix of finite-order lag polynomials. α' is a (n-k) x n matrix whose rows correspond to the co-integrating vectors; $\alpha' x_{t-1}$ represents the stationary error-correction terms. d is a n x (n-k) matrix of constants. A(L) may be obtained

from the adjoint matrix of the moving average lag polynomial C(L). Specifically, if C(L) is lower block-triangular as in the augmented system (4), then A(L) will inherit a corresponding block of zero restrictions in (5). Moreover, since all the regressors in (5) are stationary random variates, standard tools of statistical inference can be applied. Accordingly, tests of money-income causality can be straightforwardly applied while reflecting at the same time the underlying dynamics in (1). Section 4 in this paper describes the results of such tests.

Note that the disturbance term in the vector error-correction model (5) is the only observable shock term, and corresponds to the innovation in the moving average representation (2). Equation (3) demonstrates that a co-integrated model of the form (2) and (5) can be obtained from a common trends model of the form (1). Under such a construction, the innovation ε_t in general depends upon both the permanent and transitory shocks and hence possesses no natural interpretation. In testing for causality, no such interpretation is necessary, and no assumptions about the correlation between permanent and transitory disturbances is necessary at this point.

Similarly, the trend components may also be extracted from the model (1) without specifying the relationship between the permanent and transitory shocks. Using an argument analogous to that underlying equation (3),

$$W_t \equiv (H'H)^{-1}H'x_t = (H'H)^{-1}H'\delta + \tau_t + (H'H)^{-1}H'B(L)a_t.$$

Since the k-element vector W_t contains the random walks τ_t, its long-run behavior should ultimately be dominated by the nonstationary component. Accordingly, the limiting long-run forecasts of W_t based on information set I_t, $\hat{W}_{t+k}|I_t$, converge to the common trends (Stock and Watson (1987b), Watson (1986)):

$$\lim_{k\to\infty}(\hat{W}_{t+k}|I_t - \delta^* k) = \hat{\tau}_{t+k}|I_t = \hat{\tau}_t|I_t,$$

where $\delta^* = (H'H)^{-1}H'\delta$. The first differences of the extracted trends provide estimates of the permanent innovations, and the correlations between these innovations and contemporaneous output growth are reported in the first half of section 5.

On the other hand, when the intent is to uncover the sources of transitory shocks, it does become necessary to provide some interpretation of the observed disturbance in terms of the underlying unobserved transitory innovations. This corresponds to assuming that the permanent innovations η_t are linear combinations of the transitory innovations ε_t (see King, Plosser, Stock and Watson (1987)). This restriction also permits a one-to-one connection between the common trends model (1) and the co-integration model (2) and (5), with the observed disturbances representing transitory shocks. The second half of section 5 investigates the mutual correlations between these shocks.

The specification of the state vector in (5), and of the underlying propagation mechanisms in (1), are consistent (at least in spirit) with more detailed, specific analyses of RBC models. As such, the subsequent empirical analysis of (5) speaks directly to the points raised by King (1986) concerning tests of RBC models. Nonetheless, several

qualifications should be raised. On the one hand, potential misspecification (either of the state vector or of the dynamics) can never be dismissed, giving rise to spurious rejections of the RBC model. On the other hand, the current state of the theory remains silent as to the number of common trends, and as to the specific structure of C(L) in (2) (beyond the lack of causality from money to income). These considerations, together with the potentially high dimensionality of the real state vector x_t, suggest that the tests may have low power in detecting rejections of the model. It seems difficult, however, to resolve these issues through a priori reasoning, particularly in comparison with alternative testing methodologies for alternative paradigms. In any event, the subsequent analysis may be considered as a preliminary to more refined tests.

3. Data and Model Specification

a. Data

This section of the paper describes the data used in the subsequent analysis, and analyzes the univariate properties of the variables, as a preliminary to specification of the equation system. The data are drawn from six industrialized countries: Canada, the United States, Japan, West Germany, Italy and the United Kingdom. For each country, observations are obtained over the period January 1955 to December 1986 on an industrial production index for each of seven industries: mining and quarrying (ISIC code number 2); food, beverages and tobacco (ISIC code 31); textiles, apparel and leather (ISIC code 32); chemicals (ISIC code 35); basic metals (ISIC code 37); fabricated metal products,

machinery and equipment (ISIC code 38); electricity, gas and water (ISIC code 4). These seven variables are assumed to constitute the state vector describing the real sector of each country. Admittedly, the coverage does not extend to several potentially important economic sectors such as agriculture, construction or services. On the other hand, this specification of the state vector conforms to the Long-Plosser (1983) model, while preserving some tractability.

In the subsequent analysis, the vector of the real state variables is augmented by including (not all at the same time) four additional variables: the money supply (corresponding to the M1 definition of each country), a nominal interest rate, the monetary base ("outside" money), and the quantity of real bank debits for each country. The interest rate variable is generally a short-term rate (either a Treasury bill rate, or a call money rate), although the yield on long-term government bonds is used for Italy. The level of real bank debits is used as a proxy for the inputs of financial intermediation services, and is measured as the volume of nominal bank debits, deflated by the consumer price index. All the real and nominal variables are transformed by taking natural logarithms, and are not seasonally adjusted. Further details on the data, and their sources, are contained in the appendix.

b. Univariate tests for unit roots and trends

Since the appropriate treatment of stochastic and deterministic trends is central to the specification of the model, the univariate time series properties of the data are addressed first. Table 1 tests

for the existence of a unit root, and also for a deterministic time trend of up to second order, underlying the level of each variable, for each country. Part (a) of the table reports the results for the real variables, while part (b) reports corresponding results for the nominal and financial variables.

The test for a unit root is provided by the Dickey-Fuller (1979) τ_τ-statistic, applied to a fifteenth order autoregression for each variable, where monthly seasonal dummy variables are included.[4] For the sectoral production series, the null hypothesis of a unit root is rejected (at the five percent level) in only two out of 42 cases. The money stock, the monetary base and real bank debits also display a unit root for all countries in the sample, but the interest rate variable provides more mixed evidence. The null hypothesis of a unit root in this variable is rejected for Japan, West Germany and the U.K., but cannot be rejected for the other three countries.

Given the presence of a unit root in the autoregressive polynomial for most variables, Table 1 also tests for a drift term in the first differences of each variable (that is, for a linear time trend in the levels of each variable). The outcome of this test is reported as the "t"-statistic on the intercept from a regression of each differenced series on a constant, monthly seasonal dummies, and 15 of its own lagged values. Similarly, a test for a time trend in the first differences of each series (corresponding to a quadratic trend in the levels of the series) may be obtained from the "t"-statistic on time from a regression of each differenced series on a constant, time, monthly seasonal dummy variables and 15 of its own lagged values.

Part (a) of Table 1 provides the general impression that real sectoral production in each country is stationary in first differences, possibly around a non-zero drift and a time trend. In part (b), growth in inside money, real debits and, to a lesser extent, growth in outside money, display non-zero drifts. Additionally, money growth in the U.S. and also in Japan follows a time trend. The evidence for a unit root, with possible drift, is weakest for the interest rate series.

c. Tests for the number of common trends

The results of the previous section are not inconsistent with the notion that the time series behavior of sectoral output and money reflects both a permanent nonstationary component, as well as transitory, stationary disturbances. In turn, this characterization is compatible with a class of real business cycle models that embody the features of endogenous sustained growth (King and Rebelo (1986)). This section of the paper provides the detailed specification of the reduced form analog, as given in equation (1) above, of this class of RBC models.

In order to provide a sharper focus to the results, the analysis proceeds in the following sequence. To begin, the data are checked to see if the joint behavior of the real sectoral production series permits a decomposition into a smaller set of common stochastic trends. Tests for the number of these stochastic trends are also conducted. Next, the state vector is expanded with the inclusion of the money supply variable. The tests for the number of common trends are repeated, to see if the additional nominal state variable introduces

another common trend. Given the number of common trends, the number of co-integrating vectors in the vector error-correction model can be specified. This naive modeling strategy perhaps lacks the precision that a more explicitly model-based specification analysis might provide. Nonetheless, given the size of the state vector relative to the limited number of observations, the sequential strategy seems to me a not unreasonable way to proceed.

In order to establish the number of common trends, the Stock-Watson (1987b) q_f test-statistic is applied to the set of seven sectoral output series for each country. Briefly, the test first applies a pre-filter to the variables, in order to isolate the random walk components from the stationary components. Using the pre-filtered series, the matrix of first-order autoregression coefficients is obtained. If there are k common trends underlying the n variables, then this matrix should possess k unit eigenvalues, and (n-k) eigenvalues with modules less than unity. Hence, the hypothesis that there are k common trends can be rejected in favor of the hypothesis that there are (k-1) trends, if the k-th ordered eigenvalue has modulus less than one.

Table 2 shows the results of the Stock-Watson tests. In developing these results, the data are pre-filtered through a fourth-order vector autoregression, with seasonal dummy variables and a linear time trend (quadratic trend for Japan). The results indicate that output in the different industries for each country share a small number of fundamental stochastic trends: one each for Canada, Italy and the U.K.; two each for Japan and West Germany; and three trends for the U.S.

When the money stock is introduced into the state vector, the number of common trends generally increases by one, as Table 3 indicates. The only ambiguity arises in the case of the U.S. and Japan. On the one hand, the test indicates that the number of common trends is unchanged by including the money stock. On the other hand, a check for the importance of more trends cannot be rejected for these two countries. In drawing a conclusion, it seems best to specify a relatively small number of common trends. Aside from considerations of parsimony, an overly conservative estimate increases the number of cointegrating vectors, and thereby the number of error-correction terms in the model. Since irrelevant explanatory variables do not affect the consistency of the least squares estimator but only reduce efficiency, the subsequent causality tests err on the conservative side. To summarize, Table 4 specifies the orders of the deterministic, and common stochastic trend, components of the process for the (log-) levels of the eight element, real-monetary state vector.

4. Tests for Money-Income Causality

A striking implication of RBC analysis is that, given past values of the real state variables, lagged values of the money stock should have no marginal predictive ability. The results of the previous section permit the formulation of the statistical model (equation (5), repeated here) used to test this hypothesis:

$$(1-L)x_t = \phi + A(L)(1-L)x_{t-1} + d(\alpha' x_{t-1}) + \varepsilon_t. \qquad (5)$$

x_t is an eight-element vector, containing the sectoral output variables and the money stock (all in logarithms). ϕ includes a constant,

11 monthly seasonal dummy variables and a dummy variable with value one for months with strike activity. Since the levels of the variables contain quadratic time trends for the U.S. and Japan, their first differences in (5) also inherit a linear time trend. The autoregressive polynomial A(L) is specified to take the form $A(L) = A_1 + A_2L + A_3L^{11}$; this formulation yields a serially uncorrelated residual vector, ε_t. For each country, the number k of common trends (see Table 4)) determines the number of co-integrating vectors and thereby the number of error-correction terms (n-k) in each system. These error-correction terms, $\alpha' x_{t-1}$, are estimated as follows. Stock (1978) shows that the co-integrating vectors can be consistently estimated using ordinary least squares. In this case, the fitted values $\hat{\alpha} x_t$ are linear combinations of the smallest (n-k) principal components of x_t. Engle and Granger (1987) show that the two-step estimator, whereby the error-correction terms are first estimated, using principal component analysis, and then used as explanatory variables in estimating the vector error-correction model (5), results in consistent estimates. Moreover, the least squares estimator of the parameter covariance matrix is consistent.

Table 5 shows the results of tests for the ability of lagged money growth to help predict sectoral output growth. Note that these tests apply to causality from the differenced log-levels of money, taking as given the error-correction terms. In general, the lagged level of the money stock also appears in these error-correction terms, with non-zero weights. The relevant restrictions, however, apply to the zero

coefficients on lagged values of money growth, reflecting the block-triangular nature of the moving average representation.

Over the full sample period 1956:2-1986:12, there is strong evidence that past money growth helps to predict future sectoral outputs, given lagged values of the real variables. The only exception is for the United Kingdom, although the shorter sample period in this case suggests that the test may not be very powerful. In order to control for parameter nonstationarity and nonconstant residual variances, the tests are also replicated over the two halves of the sample period.[5] There generally continues to be strong evidence against the non-causality of real outputs with respect to transitory monetary innovations. However, the marginal significance levels of the test statistics tends to be higher, so that the noncausality hypothesis is not rejected (at the 10 percent level) for Italy in the first sub-period, nor for Japan and the United Kingdom in the second sub-period. Again, the shorter sub-periods may make these findings less reliable.[6]

In a vector autoregression applied to the first differences of U.S. industrial production, money and prices, Eichenbaum and Singleton (1987) find no evidence against noncausality from money to output. The results for the U.S. from Table 5 above are similarly obtained with differenced data, but use a different state vector, and include as explanatory variables a time trend and the error-correction terms (which enter significantly into the regressions). The findings here tend to agree more with those of Stock and Watson (1987a) and Bernanke (1986).

There are, on the other hand, at least two possible responses to the evidence of a strong causal role for money. Both concern some form of reverse causation in the money-output correlation. One response draws on Sims's finding (1980a, 1980b) that including an interest rate in the state vector reduces the predictive power of lagged money. Under this line of reasoning, changes in expected future real activity and expected future profitability affect current yields on financial assets, in turn changing the demand for (real) money balances. If the monetary authority passively adjusts the money supply to accommodate money demand, then changes in the money stock appear to lead changes in real activity, unless the expectational role of interest rate changes is accounted for. Such an argument would call for including an interest rate variable in the state vector. A second response arises from the analysis of King and Plosser (1984). Changes in expected future real activity affect the current demand for productive inputs, including the demand for financial intermediation services. If the output of such transaction services is positively correlated with the stock of "inside" money, then changes in "inside" money will also appear to lead changes in real activity (just as the output of any one sector helps to predict output in another sector). This response suggests that a more appropriate measure of monetary policy actions is the quantity of "outside" money, or the monetary base.

There is evidence, summarized in Table 6, that the introduction of the interest rate into the state vector also introduces an additional common trend. Accordingly, the revised vector error-correction model (5), applied to the nine-element state vector, includes the re-estimated

error-correction terms and first differences of the nominal interest rate (at lags one, two and 12). Tests for the incremental predictive power of the inside money stock and the interest rate are displayed in Table 7. Over the full sample period, there is little evidence to alter the earlier findings of a strong causal role for inside money for all countries other than the United Kingdom. Indeed, transitory interest rate innovations appear to be important only in the United States and Japan. When the sample period is split into two halves, the causal role of money is still evident, albeit at higher marginal significance levels. Interestingly, the clearest evidence of a causally prior role for transitory interest rate innovations is from the United States.[7]

Using an innovation accounting approach, Sims (1980b) finds that interest rate surprises have a significant negative impact on industrial production in a system containing the levels of output, money, prices and the interest rate. The role of monetary innovations is correspondingly diminished. His sample includes data for the U.S., West Germany, France and the U.K. On the other hand, Bernanke (1986) finds (as do Stock and Watson (1987a)) that, in the U.S., first-differencing and detrending the data maintains the predictive power of money, even with an interest rate included. The results in Table 7 confirm the findings of Bernanke (1986) and Stock and Watson (1987a) for the U.S. and also for Japan (the data for which are also detrended first differences). In addition, since the systems for the other four countries do not include a time trend, the persistent finding of

causality from money to output cannot be an artifact solely due to inappropriate detrending.

Further evidence on the nature of the money-income correlation is provided in Table 8. Here the causality tests are applied to a state vector which includes, in addition to the sectoral output series, the monetary base.[8] For comparative purposes, the results for the system with inside money are also replicated over the same sample period. If the outside money measure is used, the null hypothesis of non-causality cannot be rejected for Canada, the United States, Japan and the United Kingdom. The data for Germany reject the noncausality hypothesis, while the noncausality hypothesis is rejected for Italy at the 10 percent level (but not at the five percent level). Much of the correlation between money and real activity can thus be attributed to the "inside" component of the money stock, as the analysis of King and Plosser (1984) suggests.

A more direct test of the King-Plosser (1984) hypothesis on reverse causation is made possible by using a measure of the level of intermediation services. Here, this variable is proxied by the volume of real bank debits, which is added to the vector of sectoral outputs. If these real variables jointly summarize the state of the economy, then innovations in outside money or in the nominal interest rate should not help to predict future real activity. Table 9 replicates the tests of causality for the 10-variable vector error-correction model comprising the sectoral outputs, real bank debits, nominal base money and the nominal interest rate. Since data on real debits are unavailable for Italy, results are reported for only five countries.

The results in Table 9 help to confirm the importance of real transaction services. With the exceptions of Canada and possibly West Germany, real bank debits help to predict output in the other sectors; indeed, the test statistics on this variable in Table 9 are at least as large as those for nominal money over the same sample period (see the last column in Table 8). Nonetheless, the joint predictive power of the nominal base and the interest rate is still significant for the U.S., Japan and West Germany. At least part of this outcome, however, is due to the predictive ability of the interest rate, so that the monetary base is individually significant only for Japan and West Germany, but not for the U.S.

These findings together suggest one possible interpretation which is not unfavorable to real business cycle theory. Specifically, the level of produced financial intermediation inputs (as proxied by real debits) and an expectational variable such as the interest rate lead real outputs. Given these two variables, the role of monetary policy (as proxied by the monetary base) is substantially reduced. An alternative, less charitable, interpretation has been noted by McCallum (1983, 1986). In particular, since monetary policy (at least in the U.S.) has been effected through an interest rate instrument, innovations in the interest rate may equally well be interpreted as monetary surprises; similarly, the monetary base is not entirely appropriate as a measure of exogenous policy decisions. At the same time, the results of this paper may shed further light on some of McCallum's criticisms. In particular, the use of an expanded vector of sectoral outputs may help to reduce the possibility of omitting relevant state variables.

Formal statistical methods are used to distinguish between stochastic and deterministic trends, thereby providing a firmer basis for the statistical properties of the causality test results. Finally, extending the sample to other countries serves as a check on the robustness of the conclusions from existing tests of RBC models using U.S. data.

5. The Role of Permanent and Transitory Innovations

The proposition that monetary impulses are unimportant sources of cyclical disturbances is perhaps an exaggerated characterization of the real business cycle approach. In particular, a causal role for monetary innovations does not preclude the possibility either that permanent shocks account for a substantial portion of output variability, or that transitory shocks are mutually uncorrelated and disaggregative in nature. This section of the paper investigates these two propositions of the RBC approach, using the eight-variable system comprising sectoral output growth rates and growth in the nominal money supply. Money supply growth is included since the results of the previous section suggest that this variable has significant ability to predict future output, either because of non-neutral effects or because of reverse causation.

One way to analyze the role of permanent disturbances is to uncover the moving average representation of the system, and to track the effects of unit innovations to each of the stochastic trends, as in King, Plosser, Stock and Watson (1987). They are able to do so, however, in part because the stylized growth model in that paper delivers clues as to the nature of each trend component, and also as to

the long-run response of the system to each trend. In this paper, however, the expanded number of state variables (and trend components) significantly complicates any interpretation of the identities of the trends, their long-run effects, or their mutual correlations. At the same time, however, the expanded state vector follows from Long and Plosser (1982), and allows removal of the assumption in King, Plosser, Stock and Watson (1987) that there is only one disturbance to the system. It therefore seems unadvisable in the present context to place too much emphasis on accounting for innovations in the moving average representation of the system. Instead, the objectives of this section are more limited: namely, to gauge the quantitative importance of permanent innovations relative to the monthly variability of growth in output and money, and to determine whether the transitory innovations reflect a smaller set of aggregate disturbances. These restricted objectives limit the need to make perhaps implausible assumptions, but preclude an analysis of the system's response to hypothetical perturbations.

Alternatively, a simpler measure of the importance of fluctuations in the permanent component is given by the squared multiple correlation coefficient between growth in each industry's output and the set of contemporaneous innovations to all the stochastic trends within each country. This statistic measures the portion of variation in sectoral output accounted for by the permanent innovations,[9] and sidesteps the difficulties of attaching specific interpretations to each trend component and decomposing their correlation structure. The procedure for extracting the permanent nonstationary components is as

described in Stock and Watson (1987b). If, for each country, there are k permanent components identified by the test for common trends, then the k largest principal components of the eight-variable dataset are extracted. Each principal component represents a linear combination of the underlying nonstationary components. Hence, the long-run forecasts generated from a vector autoregression fitted to these principal components should converge to the underlying random walk processes, and yield estimates of the permanent innovations. This trend-extraction procedure does not rely on any assumption concerning the correlation between permanent and transitory disturbances. However, the measured correlations will in general also reflect the true underlying correlation between the transitory and permanent innovations to output growth.

Table 10 displays the correlations between the permanent innovations and the growth rates of sectoral outputs and money, as well as the median correlation for each country. To facilitate comparison across countries, the correlation coefficients are adjusted for degrees of freedom. The median correlations range from 0.17 for Italy to 0.55 for the United States, so that roughly 20 to 60 percent of the monthly variability in output growth can be attributed to variations in the permanent trend. There also appear to be differences across industries in the magnitude of the correlations: the correlations tend to be higher for mining, chemicals, metals and machinery, with lower correlations for food and textiles. Money growth also appears to be only weakly correlated with innovations in the permanent component.

It is also of interest to examine the correlations between transitory shocks to different sectors. The magnitude of these correlations is pertinent in distinguishing between two interpretations of cyclical fluctuations in output. On the one hand, the transitory shocks could all reflect a small set of underlying aggregate disturbances, in which case the transitory shocks should be highly correlated and be summarized by a few latent variables. This interpretation corresponds to the usual notion of an aggregate business cycle shock. On the other hand, as Long and Plosser (1987) argue, the temporary innovations could be only weakly correlated and reflect sector-specific events, which are nonetheless propagated through other sectors through agents' consumption-investment decisions. These alternative views have different implications for the appropriateness of counter-cyclical stabilization policies.[10]

The vector error-correction model analyzed in the previous section yields a series of estimated transitory innovations for each industry. Table 11 summarizes the correlations across sectors for each country, using the root mean squared correlation between each sector's innovations and the remaining innovations. The average correlations are generally small, with the highest being 0.43. Moreover, the larger correlations tend to be concentrated in textiles, chemicals, metals and machinery, perhaps suggesting a more specific factor influencing these industries. Again, it is interesting to note that the contemporaneous correlations between money growth and the other variables are low.

A more formal analysis of these results is possible, by conducting a simple factor analysis of the correlation matrix of the transitory innovations. Two factors are specified, and panel (a) of Table 12 reports the communalities of (proportion of variance explained by) the two factors. One measure of the adequacy of a two-factor model is given by the discrepancy between the actual correlation matrix and the fitted correlation matrix; panel (b) of Table 12 reports the root mean square of the off-diagonal elements of this matrix of discrepancies.

In Table 12 (b), the root mean square discrepancy is small, so that a two-factor model appears to summarize adequately the covariation between sectoral shocks. On average, the two factors account for between 20 to 40 percent of the variance of transitory innovations. In other respects, however, Table 12 (a) confirms the impression from Table 11 that much of the explanatory power of the factor model stems from its ability to account for the correlations between the textiles, chemicals, metals and machinery sectors. This would appear to be not entirely consistent with the notion that these factors reflect aggregate sources of macroeconomic disturbances.

Finally, it might be objected that the results of this section are incompatible with the findings in the previous section concerning a causal role for lagged money supply growth. In particular, the evidence in Tables 10 to 12 suggests that money growth is only weakly contemporaneously correlated with innovations to the permanent trend components, and that transitory money shocks are also weakly contemporaneously correlated with the other sectoral disturbances. It is, however, the focus on contemporaneous correlations that distinguishes

the analysis of this section from that of the previous section, with its more extended consideration of lagged cross-correlations. In this respect, the findings of this section do not necessarily conflict with the results in section 4.

6. Summary and Conclusions

This paper addresses three questions commonly identified with the real business cycle approach: do nominal variables (specifically, the money supply) Granger-cause real outputs?; do permanent innovations account for a substantial portion of variability in output growth?; do the transitory disturbances reflect a small set of common aggregate shocks, or are there many diverse shocks? These questions are analyzed in the context of a statistical model whose state vector parallels that in the theoretical model of Long and Plosser (1982), and which explicitly considers the existence of deterministic and common stochastic trends.

The results suggest that innovations in the stock of inside money are strongly causally prior to innovations in sectoral production. This finding characterizes the experience of most of the countries in the sample (with the exception of the United Kingdom); it is robust to variations in the sample period, and to the inclusion of an interest rate in the state vector. This result is apparently in conflict with the simplest form of real business cycle model, which denies the importance of money supply shocks. The finding is all the more striking, since the parameterization of the model and the limited number of observations suggests that the test should, if anything, lack the statistical power to reject the null hypothesis of noncausality.

On the other hand, the incremental predictive ability of innovations in the monetary base is much lower. For countries other than the U.S. and Japan, interest rate innovations also appear to have little incremental predictive ability. Indeed, the evidence, while mixed, is not incompatible with a direct test of the "reverse causation" argument advanced by King and Plosser (1984) and Sims (1980a,b). Real bank debits (a proxy for financial intermediation services) and financial asset yields help to predict future real activity while outside money plays a much smaller role.

In uncovering the sources of variability in sectoral output growth rates, an analysis of contemporaneous correlations provides a more favorable verdict concerning the other two propositions of the real business cycle approach. Specifically, between 20 to 60 percent of the monthly variability in sectoral output growth is associated with variability in the innovations to the permanent trend components. Across sectors, the transitory innovations are weakly correlated; at the very least, it is hard to reconcile the observed contemporaneous co-movements in output with the presence of a small set of common aggregate factors.

Taken together, the results complement, and further emphasize, a point made in King and Plosser (1987) concerning the substantively important features of business cycle behavior. Specifically, with respect to the sources of output fluctuations, the present analysis suggests an important role for innovations to the permanent, long-run trends in outputs. In comparison, monetary innovations, while having only weak contemporaneous effects, are propagated over time to yield

an important impact. It seems reasonable to hope that a common set of underlying propagation mechanisms, perhaps involving substitution across time and goods in consumption and production, can account for the effects of, on the one hand, shocks in the secular growth component, and on the other hand, disaggregate shocks in the cyclical component. While the link from money to output, especially, does not seem to be well-understood, the role of intermediate inputs of financial intermediation services offers interesting insights which are not entirely at odds with the data.

Footnotes

[1]The introduction of nominal variables into the system is, admittedly, ad hoc, but should serve as a useful starting point. Eichenbaum and Singleton (1987) present an alternative, worked-out real-nominal model.

[2]The block-triangular nature of the moving average matrix polynomial in (3) restricts the nature of the permanent and transitory components of the real variables, in the following manner. Integrating (4) yields (with initial values set to zero)

$$\begin{pmatrix} x_t \\ y_t \end{pmatrix} = \begin{pmatrix} C_{xx}(1) & C_{xy}(1) \\ C_{yx}(1) & C_{yy}(1) \end{pmatrix} \begin{pmatrix} \nu_{xt} \\ \nu_{yt} \end{pmatrix} + \begin{pmatrix} C^*_{xx}(L) & C^*_{xy}(L) \\ C^*_{yx}(L) & C^*_{yy}(L) \end{pmatrix} \begin{pmatrix} \varepsilon_{xt} \\ \varepsilon_{yt} \end{pmatrix}$$

where $\nu_{xt} = \sum_{s=1}^{t} \varepsilon_{xs}$, $\nu_{yt} = \sum_{s=1}^{t} \varepsilon_{ys}$, and $C^*_{ij}(L) = [C_{ij}(L)-C_{ij}(1)](1-L)^{-1}$ for $i,j = x,y$. If $C_{xy}(L) = 0$, then $C_{xy}(1) = C^*_{xy}(L) = 0$, so that the permanent, nonstationary part of x_t is given by $C_{xx}(1)\nu_{xt}$ while the transitory, mean-reverting part is given by $C^*_{xx}(L)a_{xt}$. Neither the permanent nor transitory components of the real variables are affected by nominal innovations.

[3]In the subsequent empirical analysis, ϕ also includes a set of seasonal dummy variables, a dummy variable for months with strike activity, and may also include a time trend. The co-integration properties of the data also impose restrictions on the seasonal and time trend components across variables. To avoid clouding the outcomes of the causality tests, however, these restrictions are not imposed. This sacrifice of efficiency would understate rejections of non-causality.

[4]The choice of a long lag length is motivated by the results of Schwert (1987), who finds that tests for unit roots are sensitive to the existence of a moving average term in the time series representation.

[5]The error-correction terms are also re-estimated over the two sub-periods.

[6]When the U.S. sample is split at October 1979, the chi-squared test for causality from money to output is 33.20 (with marginal significance level of four percent) over the first sub-period. The corresponding test statistic, based on data from 1979:10 to 1986:12, is 35.78 (with marginal significance level of two percent).

[7]Corresponding results for the U.S. sample, using data before and after October 1979, are as follows. In the pre-October 1979 period, the causality test statistic for money is 31.37 (marginal significance level of 0.07), for the interest rate 41.88 (0.00). In the post-October 1979 sub-period, the causality test statistic for money is 32.15 (0.06), and for the interest rate 36.11 (0.02).

[8]As in the analysis above, all variables enter at lags one, two and 12. The error-correction terms are re-estimated with the outside money measure. In addition, the specification of the deterministic component in the vector error-correction model is revised, in line with the results for the monetary base in Table 1(b). Specifically, the system for Canada and Japan include a linear time trend, while the other countries' systems do not.

[9]In the univariate case, the importance of the trend component can be measured by the ratio of the variance of the random walk innovations to the variance of output growth. Estimates of this variance ratio are reported in Campbell and Mankiw (1987), Clark (1987) and Cochrane (1986), using quarterly data for the U.S. The correlation coefficient reported here extends this variance ratio to the case of several stochastic trends. Moreover, by exploiting the co-integration properties of a set of variables using monthly, seasonally unadjusted data, the underlying permanent components should be more precisely measured.

[10]Blanchard and Watson (1986) pose this question as "is there only one source of shocks to the economy, or are there many?" Their results indicate that there are four, equally important, aggregate sources of shocks. On the other hand, they work with aggregate variables over a quarterly time-interval. Using data similar to that in this paper (but at quarterly and annual intervals), Stockman (1988) finds that both industry-specific and nation-specific shocks are empirically important. Norrbin and Schlagenhauf (1988) analyze disaggregate unemployment data and verify the importance of industry-specific factors, in addition to an aggregate factor.

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Table 1

Tests for unit root, drift and trend,
1956:8 - 1986:12

(a) Real variables: industrial production, by industry

Industry		Country					
		Canada	U.S.	Japan	West Germany	Italy	U.K.
2	Unit root	-1.92	-1.11	-2.36	-2.83	-2.32	-1.12
	Drift	-1.02	-2.33**	-3.34***	0.59	-1.51	-0.17
	Trend	-0.88	-1.27	-1.75*	-0.46	-1.91*	1.74*
31	Unit root	-0.39	-3.95**	-0.04	-1.98	-1.92	-0.64
	Drift	-1.84*	1.38	-2.74***	-3.22***	1.45	-3.41***
	Trend	-3.12***	-0.16	-3.34***	-3.51***	-1.24	-2.41**
32	Unit root	-1.64	-2.66	-1.20	-1.60	-4.90***	-2.58
	Drift	-1.43	0.22	-9.10***	1.48	1.99**	-3.19***
	Trend	-1.10	-0.55	-2.39**	-1.46	-0.72	-0.25
35	Unit root	-1.72	-1.18	-0.76	-0.48	-1.06	-1.29
	Drift	-0.47	-2.63***	-5.91***	1.50	1.09	1.06
	Trend	-0.87	-1.88*	-2.41**	-3.27***	-3.13***	-1.41
37	Unit root	-2.27	-1.58	-0.66	-1.97	-1.13	-1.77
	Drift	4.09***	3.01***	-3.85***	2.84***	1.46	0.76
	Trend	-0.24	-1.68*	-3.05***	-1.05	-1.70*	-1.10
38	Unit root	-1.95	-2.79	-1.26	-2.28	-2.19	-1.55
	Drift	2.79***	-1.01	-6.44***	-1.92*	2.58***	-0.02
	Trend	0.16	-0.25	-2.69***	-0.73	-1.30	-1.44
4	Unit root	-0.57	0.96	-0.58	-0.05	-1.01	-0.67
	Drift	3.79***	2.88***	-1.38	1.39	5.31***	3.29***
	Trend	-2.38**	-3.55***	-2.57**	-1.56	-1.81	-1.90*

Table 1 (continued)

(b) Nominal and financial variables: money stock, interest rate, monetary base and real bank debits

Variable		Country					
		Canada	U.S.	Japan	West Germany	Italy	U.K.
Money stock[a]	Unit root	-2.16	-0.04	0.04	-2.13	-2.29	-2.17
	Drift	-4.39***	-2.02**	-5.56***	-5.51***	-4.71***	-3.61***
	Trend	0.81	3.69***	-2.20**	-1.49	0.62	1.14
Interest rate	Unit root	-1.83	-1.99	-3.50***	-3.27**	-1.00	-2.62*
	Drift	-0.59	-0.37	-3.29***	-1.44	-3.06***	-0.59
	Trend	-0.59	-0.74	-0.41	0.09	-0.14	-0.60
Monetary base[b]	Unit root	-0.57	-2.34	-1.36	-1.95	-1.97	-0.92
	Drift	-1.87*	0.12	-3.79***	-1.25	2.24**	-3.31***
	Trend	-2.14**	0.94	-2.09**	-1.86*	-0.07	-0.74
Real debits[c]	Unit root	-1.27	-1.78	-2.12	-3.15	--	-1.89
	Drift	3.22***	6.58***	-3.88***	0.47	--	-5.38***
	Trend	-1.32	2.41**	0.36	-1.28	--	0.90

Notes: *denotes significant at the 10% level,
**significant at the 5% level, and
***significant at the 1% level.

All variables are expressed as natural logarithms and are not seasonally adjusted. The test for a unit root is the Dickey-Fuller t-statistic, with a time trend and seasonal dummy variables included. The test for drift and trend are the t-statistics of a regression of the differenced variable on a constant and time, respectively (with seasonal dummy variables included), and 15 of its own lagged values.

[a]The sample period for the U.K. is 1973:1 to 1986:12 for the money stock variable.

[b]The sample period for Canada, U.S., Japan, W. Germany and U.K. is 1967:1 - 1986:12, for Italy 1967:1 - 1985:10.

[c]The sample period for Canada is 1967:1 - 1985:6; for all other countries, 1967:1 - 1986:12. Data for Italy are unavailable.

Table 2

Tests for number of common trends underlying sectoral production, by country

Null hypothesis vs. alternative hypothesis

Country	1 root vs. 0 root	2 roots vs. 1 root	3 roots vs. 2 roots	4 roots vs. 3 roots
Canada	-5.19	-47.80***	-50.83***	-52.67**
U.S.	-2.29	-24.97	-24.88	-53.33**
Japan	-21.79	-19.45	-41.43*	-36.88
W. Germany	-2.10	-19.77	-40.14**	-43.98*
Italy	-4.80	-44.31***	-137.09***	-135.47***
U.K.	-1.85	-75.95***	-81.48***	-90.32***

Notes: Sample period for all countries is 1956:8 - 1986:12.

*denotes significant at the 10% level,
**at the 5% level, and
***at the 1% level.

Table 3

Tests for number of common trends underlying
sectoral production and money stock, by country

Country	Number of trends under null vs. under alternative	Test result	Number of trends under null vs. under alternative	Test result
Canada	2 vs. 1	-5.15	3 vs. 2	-61.64***
U.S.	4 vs. 3	-51.43*	5 vs. 4	-71.89
Japan	3 vs. 2	-46.02**	4 vs. 3	-42.52
W. Germany	3 vs. 2	-26.94	4 vs. 3	-46.53*
Italy	2 vs. 1	-7.63	3 vs. 2	-49.54***
U.K.	2 vs. 1	-5.49	3 vs. 2	-49.91***

Note: The sample period for the U.K. is 1973:1 - 1986:12; for all other countries 1956:8 - 1986:12.

*denotes significant at the 10% level,
**at the 5% level, and
***at the 1% level.

Table 4

Summary of specification of mixed real-monetary system (log-levels of variables)

Country	Order of deterministic trend	Number of common stochastic trends
Canada	1	2
U.S.	2	4
Japan	2	3
W. Germany	1	3
Italy	1	2
U.K.	1	2

Notes: The state vector for each country includes industrial production for each of the seven industries, and the nominal money stock. All variables are expressed as natural logarithms and are not seasonally adjusted.

Table 5

Tests of Granger-causality from money stock to sectoral production, by country and by sample period

	Sample Period		
Country	1956:2 - 1986:12	1956:2 - 1971:6	1971:7 - 1986:12
Canada	49.81 (0.00)	30.30 (0.09)	31.93 (0.06)
U.S.	38.80 (0.01)	31.76 (0.06)	30.09 (0.09)
Japan	37.68 (0.01)	44.13 (0.00)	28.72 (0.12)
W. Germany	53.81 (0.00)	31.63 (0.06)	35.15 (0.03)
Italy	32.07 (0.06)	24.41 (0.27)	47.22 (0.00)
U.K.[a]	27.63 (0.15)	32.07 (0.06)	23.14 (0.34)

Notes: The test statistic is asymptotically distributed as a chi-squared random variable with 21 degrees of freedom. Parenthesized entries report the marginal significance level of the associated test statistic.

[a]The overall sample period for the U.K. is 72:7 - 86:12; the sub-samples are 72:7 - 79:9 and 79:10 - 86:12.

Table 6

Tests for number of common trends underlying sectoral output, money stock and interest rate, by country

Country	Number of trends under null vs. under alternative	Test result
Canada	3 vs. 2	-24.35
U.S.	5 vs. 4	-54.45
Japan	4 vs. 3	-26.90
W. Germany	4 vs. 3	-40.01
Italy	3 vs. 2	-13.18
U.K.	3 vs. 2	-16.14

Notes: The sample period for the U.K. is 1973:1 - 1986:12; for all other countries, 1956:8 - 1986:12.

*denotes significant at the 10% level,
**at the 5% level, and
***at the 1% level.

Table 7

Tests of Granger-causality from money stock or interest rate to sectoral production, by country and by sample period

	1956:2 - 1986:12		1956:2 - 1971:6		1971:7 - 1986:12	
Country	Causality from money stock	Causality from interest rate	Causality from money stock	Causality from interest rate	Causality from money stock	Causality from interest rate
Canada	50.19 (0.00)	10.91 (0.96)	32.26 (0.06)	29.39 (0.10)	29.61 (0.10)	14.58 (0.84)
U.S.	35.42 (0.03)	49.24 (0.00)	34.43 (0.03)	32.84 (0.05)	28.67 (0.12)	37.47 (0.01)
Japan	39.11 (0.01)	53.49 (0.00)	38.91 (0.01)	30.50 (0.08)	30.03 (0.09)	24.20 (0.28)
W. Germany	45.39 (0.00)	28.96 (0.12)	28.84 (0.12)	26.26 (0.20)	32.99 (0.05)	31.17 (0.07)
Italy	32.63 (0.05)	14.56 (0.84)	23.27 (0.33)	14.97 (0.82)	41.56 (0.00)	16.62 (0.73)
U.K.	21.73 (0.42)	16.01 (0.77)	30.78 (0.08)	17.82 (0.66)	22.90 (0.35)	16.49 (0.74)

Notes: Each test statistic for the hypothesis that sectoral production can be predicted by lagged values of either the money stock or the interest rate is asymptotically distributed as a chi-squared random variable with 21 degrees of freedom. The associated marginal significance level is reported in parentheses.

[a]The overall sample period for the U.K. is 1972:7 - 1986:12; the sub-samples are 1972:7 - 1979:9 and 1979:10 - 1986:12.

Table 8

Tests of Granger-causality from monetary base
or money stock, by country

Country	Sample period	Causality from monetary base	Causality from money stock
Canada	68:2 - 86:12	27.23 (0.16)	42.32 (0.00)
U.S.	68:2 - 86:12	20.84 (0.47)	37.14 (0.02)
Japan	68:2 - 86:12	24.65 (0.26)	30.55 (0.08)
W. Germany	68:2 - 86:12	35.24 (0.03)	35.68 (0.02)
Italy	68:2 - 85:10	30.58 (0.08)	38.33 (0.01)
U.K.	72:7 - 82:10	21.65 (0.42)	27.31 (0.16)

Notes: For each country, the third column reports the test of the hypothesis that in a system including sectoral production and the monetary base, lagged values of the base predict sectorial production. The test is carried out over the sample period reported in the second column. The fourth column reports the test of the hypothesis that, over the same sample period, in a system including sectoral production and the money stock, lagged values of the money stock predict sectoral production. In large samples, each test statistic follows a chi-squared distribution with 21 degrees of freedom; marginal significance levels are reported in parentheses.

Table 9

Tests of Granger-causality from real bank debits, monetary base or interest rate

	Causality from:			
Country	Real bank debits	Base and interest rate	Base	Interest rate
Canada	19.72 (0.54)	42.54 (0.45)	26.45 (0.19)	16.68 (0.73)
U.S.	43.26 (0.00)	60.94 (0.03)	19.57 (0.55)	39.51 (0.01)
Japan	50.21 (0.00)	56.67 (0.06)	30.58 (0.08)	34.33 (0.03)
W. Germany	29.25 (0.11)	57.70 (0.05)	34.96 (0.03)	21.33 (0.44)
U.K.	50.65 (0.00)	39.65 (0.57)	26.52 (0.19)	12.76 (0.92)

Notes: For each country, the vector error-correction model includes lagged growth rates of sectoral output, real bank debits, monetary base and the interest rate. All variables enter at lags one, two and 12. Each statistic represents a test for the joint significance of each set of variables. In large samples, the test statistic follows a chi-squared distribution with 21 degrees of freedom; marginal significance levels are in parentheses. For the U.S., Japan, W. Germany and the U.K. the sample period is 1968:2 - 1986:12. For Canada, the sample period is 1968:2 - 1985:6.

Table 10

Squared multiple correlation coefficients between permanent innovations and sectoral output growth, 1956:2 - 1986:12

	Country					
Industry	Canada	U.S.	Japan	W. Germany	Italy	U.K.
Mining	0.51	0.68	0.44	0.31	0.13	0.21
Food	0.15	0.09	0.18	0.33	0.03	0.08
Textiles	0.18	0.51	0.20	0.45	0.20	0.34
Chemicals	0.24	0.29	0.38	0.54	0.14	0.26
Metals	0.43	0.68	0.36	0.63	0.20	0.57
Machinery	0.47	0.68	0.45	0.50	0.27	0.46
Electricity	0.08	0.59	0.12	0.74	0.66	0.02
Money	0.13	0.04	0.32	0.05	0.03	-0.01
Median	0.21	0.55	0.34	0.48	0.17	0.24

Table 11

Root mean squared correlation coefficients between growth rates in sectoral output, 1956:2 - 1986:12

	Country					
Industry	Canada	U.S.	Japan	W. Germany	Italy	U.K.
Mining	0.14	0.18	0.14	0.18	0.23	0.21
Food	0.19	0.18	0.11	0.32	0.25	0.26
Textiles	0.25	0.23	0.20	0.42	0.41	0.38
Chemicals	0.24	0.22	0.25	0.39	0.39	0.39
Metals	0.16	0.19	0.21	0.34	0.43	0.33
Machinery	0.21	0.20	0.20	0.43	0.43	0.42
Electricity	0.08	0.06	0.13	0.25	0.12	0.15
Money	0.09	0.10	0.05	0.06	0.08	0.09

Table 12

Estimates for two-factor model of growth rates in sectoral output and money, 1956:2 - 1986:12

(a) Communalities

	Country					
Industry	Canada	U.S.	Japan	W. Germany	Italy	U.K.
Mining	0.10	0.48	0.33	0.38	0.17	0.31
Food	0.30	0.18	0.07	0.33	0.18	0.24
Textiles	0.44	0.44	0.27	0.61	0.56	0.54
Chemicals	0.41	0.32	0.43	0.48	0.50	0.56
Metals	0.13	0.41	0.29	0.39	0.65	0.37
Machinery	0.24	0.27	0.22	0.69	0.64	0.73
Electricity	0.07	0.01	0.09	0.17	0.40	0.39
Money	0.16	0.04	0.07	0.05	0.10	0.02
Mean	0.23	0.27	0.22	0.39	0.40	0.40

(b) Root mean squared difference between actual and fitted correlations

	Country					
Industry	Canada	U.S.	Japan	W. Germany	Italy	U.K.
Mining	0.05	0.05	0.02	0.02	0.05	0.06
Food	0.04	0.06	0.03	0.05	0.05	0.07
Textiles	0.02	0.04	0.03	0.03	0.03	0.04
Chemicals	0.03	0.04	0.03	0.03	0.03	0.06
Metals	0.04	0.04	0.02	0.03	0.02	0.07
Machinery	0.05	0.03	0.05	0.03	0.03	0.04
Electricity	0.05	0.05	0.04	0.06	0.02	0.03
Money	0.05	0.06	0.02	0.04	0.03	0.05

Data Appendix

a. Sectoral production

For each of the seven industries listed in the text, output is measured as an index of industrial production. The sources are Industrial Production (OECD): Historical Statistics (1955-1971), Historical Statistics (1960-1975), supplemented by Indicators of Industrial Activity (OECD) tape. Each series is spliced to a common index of 1970: 100. For all countries, the coverage is from January 1955 to December 1986.

b. Money supply

These series are from Main Economic Indicators (OECD): Historical Statistics (1955-71), Historical Statistics (1964-1983). The data are as of the end of the month, corresponding to the M1 definition of each country. For the U.S., West Germany and Italy, changes in the definition of the series required splicing of overlapping segments of data. The coverage for countries other than the U.K. is from January 1955 to December 1986. Monthly data for the U.K. begin in June 1971.

c. Interest rate

Sources for this series are Main Economic Indicators (OECD): Historical Statistics (1955-71), Historical Statistics (1964-1983). The data are percent per annum, as of the end of the month. For Canada, U.S. and the U.K., the variable is measured as the three-month Treasury bill rate. For Japan and West Germany, the call money rate is used, and the yield on long-term government bonds is used for Italy. A change in the definition of the series required splicing the German data. For all countries, the coverage is from January 1955 to December 1986.

d. Monetary base

This variable is taken from the "Reserve Money" item in International Financial Statistics (IMF), for all countries other than the U.K. Data for this series from January 1967 are reprinted in International Financial Statistics Supplement on Money (1983), and are updated until December 1986 from monthly issues of International Financial Statistics. Due to a gap in the series, the data for Italy stop at October 1985. For the United Kingdom, the source is Bank of England Quarterly Bulletin (March 1981).

e. Real bank debits

Data on nominal bank debits are from International Financial Statistics (IMF) and extend from January 1967 to December 1986, except for Canada. The series for Canada stops at June 1985. The consumer price index is used as the deflator; data on this series are from OECD Historical Statistics (1955-71, 1964-83).

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